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Gold as an Inflation Hedge in a Time-Varying Coefficient Framework

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Gold as an Inflation Hedge in a Time-Varying Coefficient Framework

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Joscha Beckmann and Robert Czudaj¹

Gold as an Inflation Hedge in a Time-Varying Coefficient Framework

Abstract

This study analyzes the question whether gold provides the ability of hedging against inflation from a new perspective. Using data for four major economies, namely the USA, the UK, the Euro Area, and Japan, we allow for nonlinearity and discriminate between long-run and time-varying short-run dynamics. Thus, we conduct a Markov-switching vector error correction model (MS-VECM) approach for a sample period ranging from January 1970 to December 2011. Our main findings are threefold: First, we show that gold is partially able to hedge future inflation in the long-run and this ability is stronger for the USA and the UK compared to Japan and the Euro Area. In addition, the adjustment of the general price level is characterized by regime-dependence, implying that the usefulness of gold as an inflation hedge for investors crucially depends on the time horizon. Finally, one regime approximately accounts for times of turbulences while the other roughly corresponds to 'normal times'.

JEL Classification: C32, E31, E44

Keywords: Cointegration; gold price; inflation hedge; Markov-switching error correction; price level

August 2012

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1 Introduction

Although gold is no longer a central cornerstone of the international monetary system since the breakdown of Bretton Woods, it still attracts considerable interest from both investors and researchers. Owing to the increasing complexity of financial markets, diversifying a portfolio through hedging has become increasingly important. In a nutshell, gold may act as a hedge or 'safe haven' asset for portfolio investors.¹ The majority of academic research has focused on the first property since a widespread consensus states that the price of gold reflects inflation expectations. One reason is that commodity prices are generally considered to be able to incorporate new information faster than consumer prices (Mahdavi and Zhou, 1997). Gold prices seem to be appropriate regarding the reflection of inflation expectations since, in contrast to many other commodities, gold is durable, relatively transportable, universally acceptable and easily authenticated (Worthington and Pahlavani, 2007). From a theoretical point of view, an increase in expected inflation will force investors to buy gold, either to hedge against the expected decline in the value of money or to speculate due to the associated rise of the gold price. This generates a purchasing pressure which yields to an immediately rising price of gold in time of the upward revision in inflation expectations. Thus, changes in expected inflation will cause changes in the price of gold and investors with knowledge regarding future inflation have the ability to gain excess revenues by purchasing and selling gold in spot and futures markets in anticipation of prospective market adjustments. Therefore, the gold price acts as a leading indicator of the level of inflation and hence, gold could be used to hedge against future inflation. This causality pattern has recently been questioned by Blose (2010) who argues that the costs of carrying gold are also affected through changes in interest rates by expected inflation. If those costs offset revenues from speculation, the price of gold is not affected by changing inflation expectations.²

To understand the main characteristics of gold prices, a brief description of supply and demand factors is necessary. According to Baur and McDermott (2010) and the World Gold Council the supply of gold, such as the supply of other commodities, is relatively inelastic, owing to the difficult extraction process and the tedious establishment of new mines.³ Central banks also still keep maintaining passive stocks of gold, independently of the patterns of the real price of gold (Aizenman and Inoue, 2012). Hence, the gold supply remains relatively stable while the gold demand is rapidly changing in response to global economic occurrences.

¹Gold has also retained its unique status in central bank portfolios (Aizenman and Inoue, 2012).

²Blose (2010) labels this view as the carrying costs hypothesis while he corresponds to the expected inflation hypothesis as the opposite view.

³Although the gold supply is inelastic to prices it has not been constant over the whole sample period under observation. It actually peaked in 2001.

Following Baur and McDermott (2010) the total demand for gold can be separated in three categories: demand for jewelry, demand for industrial and dental, and investment demand. While demand for jewelry as well as for industrial and dental is largely determined by consumer spending power and thus, attached to the business cycle, the investment demand for gold could act counter-cyclical due to an increased demand for gold in times of global crises and recessions. Ghosh et al. (2004) merely divide the total gold demand in two categories, namely the 'use demand' which combines the first two categories mentioned above and the 'asset demand', where gold is also used by governments, fund managers and individuals as an investment to hedge inflation and other forms of uncertainty.

Although gold may act as an inflation hedge in the long-run, the price for gold also displays significant short-run price volatility (Aggarwal, 1992). Hence, cointegration techniques seem to be the appropriate tools to examine the relationship between consumer prices and the price for gold. Those frameworks are based on the idea of a stable long-run equilibrium relation between the prices for goods and services and for gold while the variables are allowed to depart from their equilibrium path in the short-run due to random shocks in the spirit of Engle and Granger (1987). The adjustment coefficients correspond to the speed of correction of short-run deviations from the long-run equilibrium. In this vein, many previous studies have focused on the inflation hedging ability of gold by partly applying conventional cointegration frameworks (Kolluri, 1981; Moore, 1990; Laurent, 1994; Chappell and Dowd, 1997; Mahdavi and Zhou, 1997; Harmston, 1998; Ghosh et al., 2004; Levin and Wright, 2006). However, the relationship between consumer prices and gold prices has undergone several structural changes since the early seventies: the breakdown of Bretton Woods in 1973, two major oil price shocks in 1973 and 1979/80, the collapse of the USSR in 1991 and the subsequent democratization of Russia as one of the top producer of gold worldwide, the burst of the 'dot-com bubble' in 2001 and the recent financial and economic crises that started in 2007.⁴ Hence, due to the frequent changes in the equilibrium relationships the parameter constancy assumption of the traditional VECM methodology seems to be too restrictive. Worthington and Pahlavani (2007) argue that ignoring such substantial changes could yield to the inability of proving the existence of a stable long-run relationship between consumer prices and the price for gold.

However, to the best of our knowledge the only study which considers the possibility of non-linearity in the inflation hedging relation is provided by Wang et al. (2011).⁵ They argue that

⁴In addition, gold price controls were not completely phased out in the U.S. until the mid 1970's.

⁵Kyrtsou and Labys (2006) also show that the nature of the dependence and the causal relation between U.S. inflation and commodity prices can be characterized by nonlinearity using nonlinear Granger causality

the existence of transaction costs and the above mentioned business cycle dependence of the gold demand possibly results in a nonlinear relationship between consumer prices and gold prices. For this reason, they account for nonlinearity based on the threshold cointegration framework developed by Enders and Siklos (2001). From a methodological point of view, such a framework adequately describes the dynamics if the data is primarily generated by market forces since the regime variable is assumed to be endogenous (Balke and Fomby, 1997; Enders and Siklos, 2001; Lo and Zivot, 2001; Hansen and Seo, 2002). However, if gold acts as an inflation hedge, exogenous factors such as major economic shocks or changes in economic policy may have an influence on the path of the gold prices. In that case, a Markov-switching vector error correction model (MS-VECM), where the unobservable state follows an exogenous stochastic process, is more suitable compared to a conventional threshold model (Ihle and von Cramon-Taubadel, 2008). This is also reasonable from an economic point of view since choosing a specific transition variable may also indicate a certain causality since the adjustment pattern depends on the chosen threshold.

Hence, the rationale for this study is to contribute to the literature by adopting a MS-VECM approach when analyzing whether gold provides the ability of hedging inflation in four major countries, namely the USA, the UK, the Euro Area, and Japan. We allow for nonlinearity and discriminate between long-run and time-varying short-run dynamics. Strictly speaking, the benefit of our framework is that it allows identifying the potentially latent regimes in the data and helps to specify the nonlinear dynamics between the variables adequately. More precisely, we are able to provide a measure of usefulness for an inflation hedge over different time periods by analyzing the time-varying price adjustment for different subperiods.

The remainder of this paper is organized as follows. The following section provides a brief summary of previous empirical findings. Section 3 describes our data as well as our empirical framework. The results are presented and analyzed in section 4. Section 5 concludes.

2 Literature review

The dynamics of the price for gold has been focused by academic literature in many respects over the last decades. The first literature strand consists of studies which address impacts of macroeconomic variables such as exchange rates, interest rates, or output on the price for gold. Studies of this kind have been provided by Sherman (1982, 1983), Ariovich (1983), Fortune (1987), Dooley et al. (1995), Sjaastad and Scacciallani (1996), Lucey et al. (2006),

and Wang and Lee (2011). Focusing on a different topic, Koutsoyiannis (1983), Diba and Grossman (1984), Baker and van Tassel (1985), and Pindyck (1993) deal with the ability to forecast the gold price. Generally speaking, evidence for different relationships and causalities including the price of gold has been provided. However, a detailed description of the corresponding results is beyond the scope of this paper.

Focusing on the recent increase in gold prices, Bialkowski et al. (2011) test whether a speculative bubble in the price of gold exists by applying a Markov regime-switching augmented Dickey-Fuller test and conclude that the high current gold price seems to be fundamentally justified. Tests of the market efficiency hypothesis for gold markets have been undertaken by Tschöegl (1980), Solt and Swanson (1981), Ho (1985), Basu and Clouse (1993), and Smith (2002). Another frequently analyzed research topic is the capability of gold as a hedge or a 'safe haven' with regard to financial assets such as bonds or stocks.⁶ Such a question has been tackled by Sherman (1986), Jaffe (1989), Chua et al. (1990), Ciner (2001), Michaud et al. (2006), Hillier et al. (2006), McCown and Zimmerman (2006), Baur and Lucey (2010), Baur and McDermott (2010), and Ciner et al. (2010). Capié et al. (2005) examine whether gold could be a hedge against fluctuations of the U.S. dollar. The overall results are mixed with recent results suggesting that gold is only able to provide a 'safe haven' function in the very short-run (Baur and Lucey, 2010).⁷

Finally, studies which are most closely related to our topic have analyzed the linkages between inflation and the price for gold. Kolluri (1981), Moore (1990), Laurent (1994), Chappell and Dowd (1997), Mahdavi and Zhou (1997), Harmston (1998), Ghosh et al. (2004), Levin and Wright (2006), Worthington and Pahlavani (2007) and Wang et al. (2011) examine the inflation hedge effectiveness of gold by focusing on the short-run and the long-run relationship between the general price level and the price for gold. It is worth mentioning that studies which focus on 'safe haven' aspects mostly adopt daily instead of monthly data. The results which will briefly be described in the following are not clear-cut and vary depending on the sample period and the country under observation. In an early study, Mahdavi and Zhou (1997) tackle the question whether gold and other commodity prices are leading indicators of inflation based on the estimation of a conventional VECM for the USA using the Johansen (1988, 1991) framework for a quarterly sample period that ranges from 1970 to 1994

⁶Baur and Lucey (2010) define a hedge and a 'safe haven' as an asset that is uncorrelated or negatively correlated with another asset or portfolio on average and only in times of market stress or turmoil, respectively.

⁷Aizenman and Inoue (2012) also point out that central bank's gold position signals economic might, and that gold retains the stature of a 'safe haven' asset at times of global turbulence.

and conclude that consumer prices and the price of gold are not cointegrated. Thus, their outcomes are in line with the findings of Garner (1995) and Cecchetti et al. (2000) who were unable to provide empirical evidence for the usefulness of the gold price as leading indicator for inflation as well.⁸ On the contrary, Laurent (1994), Harmston (1998) and Ghosh et al. (2004) detect an effective long-run inflation hedge of gold in the USA, the UK, France, Germany, and Japan. Ghosh et al. (2004) use monthly U.S. data that ranges from 1976 to 1999 and apply cointegration regression techniques. Adrangi et al. (2003) find that gold prices are positively correlated with expected inflation and conclude that a gold investment may be a reliable inflation hedge in both the short-run and the long-run. Levin and Wright (2006) also estimate a conventional VECM to analyze the short-run and long-run determinants of the gold price for the USA over a sample period from 1976 to 2005. They identify a stable long-run relationship between the gold price and the price level. Their findings also provide evidence that the change in the gold price is positively related to the change in inflation, inflation volatility, and credit risk while a negative relationship with regard to the U.S. dollar trade weighted exchange rate and the gold lease rate prevails. Using monthly data spanning the period from September 1994 to December 2005 for 14 countries (Australia, Canada, the European Union, New Zealand, Sweden, the United Kingdom, Japan, Mexico, Norway, the USA, Brazil, China, India, and Israel) Tkacz (2007) finds that the gold price contains significant information for future inflation in several countries, especially in those that have adopted formal inflation targets.

Worthington and Pahlavani (2007) also test the presence of a stable long-run relationship between the price of gold and inflation in the USA using monthly data from 1945 to 2006 and from 1973 to 2006. However, they additionally allow for instabilities when analyzing the long-run relation. In their framework, the timing of structural breaks is endogenously determined by applying the unit root testing procedure developed by Zivot and Andrews (1992) and thereafter a modified cointegration approach suggested by Saikkonen and Lütkepohl (2000a, 2000b, 2000c) is adopted. The results provide evidence in favor of a cointegrating relationship between the price of gold and inflation in both sample periods and thus, Worthington and Pahlavani (2007) conclude that a gold investment can serve as an effective inflationary hedge. Finally, Wang et al. (2011) have analyzed the short-run and long-run inflation hedging effectiveness of gold in the USA and Japan for a sample period ranging from January 1971 to January 2010 while using monthly data. They conduct the linear cointegration test proposed by Engle and Granger (1987) as well as the nonlinear threshold cointegration test

⁸Furthermore, Tully and Lucey (2007) have applied a power GARCH approach and also do not find a significant relationship between gold prices and inflation.

suggested by Enders and Siklos (2001) and show that in low momentum regimes gold is unable to hedge against inflation in both the USA and Japan, however, in high momentum regimes, a gold investment is able to hedge against inflation in the USA, and partially hedge against inflation in Japan.

Summing up the empirical evidence, the ambiguous results of previous studies as well as the provided evidence for instabilities in the relationship between gold prices and inflation verify the application of a time-varying Markov-switching approach. In our analysis, the long-run coefficients indicate the intensity of a relationship between the price for gold and the general price level. However, gold is only useful as a hedge if prices adjust to deviations from such a long-run relationship. In this case, the long-run estimates also provide a measure of effectiveness of an inflation hedge. Hence, in our framework the time-varying adjustment coefficients are able to discriminate between periods with and without hedging functions. If prices do not adjust, buying gold may not be able to shield a portfolio with respect to future price movements during a specific period.

3 Data and econometric methodology

3.1 Data

We use a monthly dataset including the price for gold denominated in U.S. Dollar, British Pound Sterling, Euro and Japanese Yen as well as the consumer price index (CPI) and the producer price index (PPI) of the USA, the UK, the Euro Area, and Japan. Using two different measures of inflation includes a robustness test for our results and also allows for different hedging abilities of gold for consumers and producers.⁹

The gold price data has been provided by the World Gold Council and covers a sample period from December 1969 to December 2011. Hence, we also include the period immediately before the breakdown of Bretton Woods. The CPI's (2005=100) and the PPI's (2000=100) for the USA, the UK, and Japan have been taken from the statistics provided by the OECD and the IMF, respectively. Both the CPI (2005=100) as well as the PPI (2005=100) for the Euro Area are taken from the ECB Statistical Data Warehouse and cover only a period ranging from February 1980 to November 2011 and from January 1981 to November 2011, respectively. The country selection is motivated by the fact that the currencies of the chosen four major economies are regarded as key currencies by the World Gold Council. As it is

⁹While most studies regarding the relationship between inflation and gold prices such as, for instance, Wang et al. (2011) use the CPI to construct a measure of inflation, Lawrence (2003) adopts the PPI.

common practice, each series is taken as natural logarithm.

In order to analyze the underlying long-run dynamics using bi-variate cointegration techniques, it is important to assure that the times series under observation are integrated of the same order. In the present context, an important question is whether some variables, in particular prices, are integrated of order two, e.g. $I(2)$. The results of the augmented Dickey-Fuller (ADF) test and the more powerful Ng-Perron MZ_α test suggest that all series may be approximated as integrated of order one, e.g. $I(1)$ (Dickey and Fuller, 1979; Ng and Perron, 2001).¹⁰

3.2 Econometric methodology

In the following we use a Markov-switching vector error correction model (MS-VECM) to examine the relationship between the price for gold (g_t) and the CPI as well as PPI (p_t) for the set of the four countries mentioned above. The concept of the MS-VECM is based on traditional state-dependent time series models developed by Hamilton (1989) and continued by Krolzig (1997). As will be described below, the parameters of a MS-VECM are designed to take a constant value in each regime and to shift discretely from one regime to the other with different switching probabilities. The switches between different states are not accounted as deterministic occurrences, but are assumed to follow an exogenous stochastic process. Consider a M -regime p th order MS-VECM, which in general allows for regime shifts in the vector of intercept terms, the autoregressive part, the long-run matrix, and the variance-covariance matrix of the errors:¹¹

$$\Delta Y_t = v(s_t) + \Gamma(L)(s_t)\Delta Y_{t-l} + \Pi(s_t)Y_{t-1} + \varepsilon_t, \quad t = 1, \dots, T, \quad (1)$$

where Δ denotes the difference operator, Y_t represents a K -dimensional vector of the observed time series, $Y_t = [p_t, g_t]'$; $v(s_t)$ denominates a K -dimensional vector of regime-dependent intercept terms, $v(s_t) = [v_1(s_t), \dots, v_K(s_t)]'$, while $\varepsilon_t = [\varepsilon_{1t}, \dots, \varepsilon_{Kt}]'$ describes a K -dimensional vector of error terms with regime-dependent variance-covariance matrix $\Sigma(s_t)$, $\varepsilon_t \sim NIID(0, \Sigma(s_t))$. The $K \times K$ matrix lag polynomial $\Gamma(L)(s_t)$ of order p denotes the state-dependent short-run dynamics of the model. The stochastic regime-generating process is assumed to be an ergodic, homogenous and irreducible first-order Markov chain with a finite number of

¹⁰We apply an auxiliary regression either with a constant or with a linear trend plus constant since a graphical inspection shows that most of the series exhibit a time dependent mean. The results of the unit root tests are available upon request.

¹¹This model is often referred to as Markov-Switching-Intercept-Autoregressive-Heteroskedastic-VECM or MSIAH-VECM. See Sarno and Valente (2006) among others.

regimes, $s_t \in \{1, \dots, M\}$, and constant transition probabilities¹²

$$p_{ij} = Pr(s_{t+1} = j | s_t = i), \quad p_{ij} > 0, \quad \sum_{j=1}^M p_{ij} = 1 \quad \forall \quad i, j \in \{1, \dots, M\}. \quad (2)$$

The first expression of equation (2) gives the probability for switching from regime i to regime j at time $t + 1$ which is independent of the history of the process. p_{ij} is the element in the i th row and the j th column of the $M \times M$ matrix of the transition probabilities P , which is usually not symmetric.

The non-stationary behavior of the series is accounted for by a reduced rank ($r < K$) restriction of the state-dependent $K \times K$ long-run level matrix $\Pi(s_t)$, which can be fragmented into two $K \times r$ matrices $\alpha(s_t)$ and β such that $\Pi(s_t) = \alpha(s_t)\beta'$. β' gives the coefficients of the variables for the r long-run relations which are assumed to be constant over the whole sample period, while $\alpha(s_t)$ contains the regime-dependent adjustment coefficients describing the reaction of each variable to disequilibria from the r long-run relations given by the r -dimensional vector $\beta'Y_{t-1}$. Thus, in our model the most interesting distinction between regimes is the speed in which deviations from the long-run equilibrium are corrected given by $\alpha(s_t)$.

First, in order to identify the rank of $\Pi(s_t)$, e.g. the number of cointegrating relations r , and to estimate the coefficients of the r cointegrating vectors in β' we rely on a framework developed by Johansen (1988, 1991). Then, conditional on these cointegrating vectors, the regime-dependent adjustment parameters $\alpha(s_t)$, intercept terms $v(s_t)$, autoregressive coefficients $\Gamma(L)(s_t)$, and variance-covariance matrix $\Sigma(s_t)$ as well as the transition probabilities are estimated using a Markov Chain Monte Carlo (MCMC) method, namely the multi-move iterative Gibbs sampling procedure proposed by Krolzig (1997).¹³ This two-step framework is adopted from the work of Saikkonen (1992) as well as Saikkonen and Luukkonen (1997)

¹²The ergodicity assumption implies a stationary unconditional probability distribution of the regimes, the homogeneity assumption defines the transition probabilities to be constant, and finally, the irreducibility assumption assures that any regime can be reached from any other regime (Ihle and von Cramon-Taubadel, 2008).

¹³The Gibbs sampling technique goes back to Geman and Geman (1984) and has been extended to univariate Markov-switching models by Albert and Chib (1993), McCulloch and Tsay (1994), and Filardo (1994). An alternative procedure is the expectation-maximization (EM) algorithm for maximum likelihood estimation which has been applied, for instance, by Sarno and Valente (2006). See Dempster et al. (1977), Hamilton (1993), Krolzig (1997), and Kim and Nelson (1999) for details. Although, the EM algorithm requires fewer computations, it does not directly provide the posterior distribution of the parameters and estimates of the variance-covariance matrix of the errors (Krolzig, 1997). To check for robustness we have used the EM algorithm as well. Our results turned out to be reliable.

who showed that the Johansen procedure provides consistent estimates for the cointegrating vectors even in the presence of regime-switching.¹⁴

4 Empirical results

With respect to the deterministic components of a VECM, Johansen (1988) distinguishes five different specifications. In this vein, an important question is whether deterministic trends are obtained in the data and whether they cancel out in the long-run relationships. Since both the price of gold as well as consumer and producer prices may include such trends, we consider two trend specifications for each model, both allowing for a linear deterministic trend in the data. The first configuration also includes a trend in the cointegration space, implying that the underlying deterministic trends of the general price level and the price for gold do not cancel out in the long-run relationship. The second setting corresponds to the case where the trends do cancel out. In this case, not a trend but an unrestricted constant is included since each of our series shows a time dependent mean. In the following, all long-run estimations for both CPI and PPI are carried out for both configurations. In terms of robustness, this proceeding assures that the detection of a long-run relationship does not depend on the specification of the deterministic components. However, since the significance of a trend in the long-run relationship can be tested, we base our analysis of the time-varying short-run dynamics on the more adequate specification.

The choice of the lag length p of each model is based on the Schwarz criterion and tests for autocorrelation and ARCH effects. According to Rahbek et al. (2002), the rank test results we gain in the following are still robust under the remaining ARCH effects in some cases. While excess kurtosis does not introduce a significant bias to the estimated cointegrating vectors the findings are sensitive to excess skewness (Juselius and MacDonald, 2004; Juselius, 2006). After introducing dummy variables to account for outliers following the methodology described in Juselius (2006), the rejection of the normality assumption is solely due to excess kurtosis, so that our results are still reliable. The corresponding statistics are available upon request.

The results of the trace test developed by Johansen (1988, 1991) are presented in Table 1 and 2 for the CPI's and PPI's, respectively. For each price measure, the null of no cointegration can significantly be rejected while the null of one cointegrating relation ($r = 1$)

¹⁴Among others Francis and Owyang (2005) followed the same methodology.

cannot.¹⁵ Merely in two cases the null of no cointegration cannot be rejected: Those are the PPI configurations for the USA and the Euro Area both with an unrestricted constant. In both cases, we do not estimate the cointegration coefficients. These findings may be simply explained by the significance of the trend within the cointegration space in the first specification. Conditional on $r = 1$, the coefficients of the cointegrating relation between the general price level and the price for gold have been estimated for all other settings and are reported in Tables 1 and 2 as well.

Table 1 and 2 about here

According to Wang et al. (2011) the long-run relationship between the price level and the gold price should be proportional if gold is able to fully hedge inflation. The cointegration vectors have been normalized on the CPI and PPI, respectively, and in each case the gold coefficients have the expected sign; however they are less than unity in magnitude. As can be seen in Table 1 and 2, respectively, most of the gold coefficients are highly significant. The only exception is the restricted trend specification for the Euro Area where the CPI is used as the price level. In that cointegrating relation, the gold price is not significant, thus we do not estimate the time-varying adjustment coefficients in this case. Overall, the gold coefficients range from 0.23 to 0.43 in case of the USA, from 0.35 to 0.45 for the UK and from 0.15 to 0.23 in case of Japan. For the Euro Area, the significant coefficients turn out to be 0.196 (CPI) and 0.055 (PPI). Although we still have to consider the time-varying adjustment coefficients it seems that gold is only partially able to hedge future inflation in the long-run and this ability is stronger for the USA and the UK compared to Japan and the Euro Area. The findings for Japan are in line with the outcome of Wang et al. (2011) and may, for instance, be explained by the inclusion of the deflation experience since the nineties where gold prices may have lost their function as a leading indicator. In case of the Eurozone, our findings can possibly be attributed to different inflation expectations across countries prior to the introduction of the Euro. Turning to the magnitudes of the long-run coefficients, the latter hardly differ between the CPI and PPI configuration for the UK. However, the CPI coefficients are always higher for Japan and the United States compared to the PPI configurations. The highest coefficient for the Euro Area is also observed for the CPI set-up. Hence, the long-run hedge against inflation provided by gold tends to be stronger for consumer prices. Overall, the general findings are remarkable robust since the verification of a long-run relationship is always independent of the chosen price measure (PPI or CPI) and

¹⁵The test statistic of the corresponding likelihood test, the so-called trace test, is given by $trace(r) = -T \sum_{i=r+1}^K \log(1 - \hat{\lambda}_i)$. Under the null of $K - r$ unit roots $\lambda_i, i = r + 1, \dots, K$, should behave like random walks and the test statistic should be small. Starting with the hypothesis of full rank, the rank is determined by using a top-bottom procedure until the null cannot be rejected (Juselius, 2006).

only differs in two cases with respect to the deterministic components incorporated into the model.

A test of the proportionality restriction based on a likelihood ratio (LR) procedure suggested by Juselius (2006) confirms the inability of gold to provide a full hedge against inflation. The test statistics which are also given in Tables 1 and 2 confirm our results since the restriction is clearly rejected for each country.¹⁶

As a next step the MS-VECM representation given in equation (1) has been estimated for each country while we allowed for two states and let each parameter switch, the intercept terms, the autoregressive components, the variance-covariance matrix of the errors and especially the adjustment parameters to deviations from the mentioned long-run relations.¹⁷ As previously mentioned, the time-varying adjustment coefficients are able to discriminate between periods with and without hedging functions since gold is not able to shield a portfolio with respect to future price movements if prices do not adjust to the long-run relation.

To estimate the MS-VECM for each setting we apply the cointegrating relations with a restricted trend given in Table 1 and 2 (*a*), respectively, as long as the trend is significant, otherwise we use the specification with an unrestricted constant (*b*) following Juselius (2006). In some cases it appeared tractable to restrict the autoregressive components not to switch between regimes according to Francis and Owyang (2005). Tables 3 and 4 provide the estimated adjustment parameters to short-run deviations of the long-run relationships given in Tables 1 and 2 that vary across regimes and Tables 3 and 4 also report the transition probabilities between states.

Table 3 and 4 about here

As can be seen in Table 3 and 4 for each country the CPI and PPI, respectively, adjusts statistically significant to deviations from the stable long-run equilibrium. Thus, it becomes evident that the general price level reacts to variations of the price for gold that let the price level drift apart from its equilibrium path. Overall, a reasonable conclusion is that gold price movements mirror changes of inflationary expectations and hence, signal the formation of future inflation. Thus, gold can effectively be used to hedge against inflation. Furthermore, it

¹⁶Additionally, we applied LR tests of exclusion for each setting which confirm the robustness of our estimated long-run relations. To conserve space we do not report the corresponding test statistics, however these could be provided upon request.

¹⁷We have also considered more than two states, but that did not seem to be feasible in our framework since at least one state did not had enough periods identified to estimate the regime switching parameters confidently.

is obvious that the price level adjusts to short-run errors either in both regimes with different speed of adjustment as in case of the UK, Japan (CPI), and the Euro Area (CPI) or just in one regime as in case of the USA, Japan (PPI), and the Euro Area (PPI). Hence, the ability to hedge according to the long-run coefficients is not continuously in the latter cases and differs for the former settings. For instance, in case of the UK the CPI corrects deviations from the long-run relation between the CPI and the price for gold by 0.7 per cent per month in the first regime and twice as fast in the second regime. Overall, it seems that the nonlinear structure of the dynamics between the general price level and the gold price is adequately accounted for by our state-dependent approach.

Moreover, the transition probabilities are highly significant for each country and show that regimes are generally persistent in each case since the probabilities of staying in a given state are relatively large. The series of smoothed probabilities also achieved by the Gibbs sampling technique and shown in Figure 1 and 2, respectively, illustrate the reconstructed incidences of the first regime over the whole sample period by inferring the probabilities of the occurrence of the unobserved states conditional on the available information in the whole dataset which allows a 'real time' classification of the regime switches (Krolzig, 2003). For regime classification Hamilton (1989) suggested the use of a probability of 0.5.

A graphical inspection of Figures 1 and 2 shows that a switching between both regimes frequently occurs at the beginning of the sample after the breakdown of Bretton Woods. Although a definite interpretation of the economic circumstances related to specific regimes is a notorious difficult task, one regime approximately seems to account for times of turbulences such as until the mid-80's or during the recent global financial crisis while the other corresponds to 'normal times' which are unaffected by major shocks due to historical events. For the United States, the probability for regime 1 for the CPI measure and regime 2 for the PPI measure is much higher between 1985 and 2000, a time period which may be labeled as 'normal times'. In both cases, prices only adjust in the latter mentioned regime. The findings for Japan suggest that the second regime is more probable during the nineties, a period during which the Japanese economy suffered from deflation and stagnation. However, the adjustment seems to be faster during the other regime although the PPI adjustment coefficient for the first regime is a borderline case in terms of significance. As previously mentioned, the price for gold may exhibit less information with respect to future price movements in a deflationary environment. For the UK, the first regime where adjustment is (slightly) weaker seems to correspond to 'normal times'. The regime pattern for the Euro Area is less clear-cut since the seventies as a period of turbulences are not included. Thus,

the adjustment pattern of the price level to the gold price depends on whether economic turbulences occur or not. However, an unambiguous time pattern across countries cannot be observed, mirroring that each economy has its own characteristics.

Figure 1 and 2 about here

Finally, the regime classification measure (RCM) suggested by Ang and Bekaert (2002) indicates a good fit of our MS-VECM for each country (see Notes of Figure 1 and 2). The RCM statistic is computed as follows

$$RCM(M) = 100M^2 \frac{1}{T_j} \sum_{t=1}^{T_j} \prod_{j=1}^M \tilde{p}_{j,t}, \quad (3)$$

where $\tilde{p}_{j,t}$ denotes the smoothed probability for regime j and provides a degree of accuracy with which a model identifies regime switching behavior over the entire sample period or a particular sub sample period. The regime variable is Bernoulli distributed and thus, the RCM corresponds to a sample estimator of its variance. It takes values between 0 and 100, with 0 representing a perfect regime classification performance and 100 denoting that the model fails to exhibit any information about the regime-dependence. For most of the countries the value of the RCM statistic is fairly low (always below 50) and thus, one could conclude that our models are adequately specified.

5 Conclusion

This study has analyzed whether gold provides the ability of hedging inflation for four major economies, namely the USA, the UK, the Euro Area, and Japan, by allowing for nonlinearity and discriminating between long-run and time-varying short-run dynamics. To the best of our knowledge our study is the first to consider a VECM approach with Markovian shifts in the adjustment parameters to short-run deviations from the long-run relations in this context. Our main findings are threefold: First, we show that gold is partially able to hedge future inflation in the long-run. This ability tends to be stronger for consumer prices in general as well as for the USA and the UK compared to Japan and the Euro Area. Secondly, the adjustment of the general price level seems to be appropriately characterized by regime-dependence. Finally, we display that one regime may account for times of turbulences and the other for 'normal times' which are unaffected by major shocks. Thus, the different adjustment pattern of the price level may depend on the occurrence of economic turbulences. Due to this finding, the price for gold should be considered when aiming to appropriately

forecast the inflation rate.¹⁸

From an investor's point of view, the effectiveness of gold as an inflation hedge crucially depends on the time horizon. Over the very long-run, gold is useful as a partial hedge since a cointegrating relationship prevails. However, during some periods where no price adjustment is observed, gold is not able to shield a portfolio. This may be illustrated by a simple example of an investor who buys gold at the beginning of a period with no adjustment and sells at the end of the corresponding period.

However, since each economy has its own characteristics, we are unable to detect an overall unambiguous time pattern across countries. An interesting topic for further research may be the analysis of gold as a hedge or 'safe haven' with respect to stock markets in a time-varying framework. The possibility to apply daily data in such an analysis may contribute to further insights with respect to the importance of gold for portfolio investors over the very short-run. Another appealing question is how does gold perform against other potential inflation hedges such as stocks, bonds, real estate, or other commodities in our framework.¹⁹

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¹⁸For instance, the gold price could be incorporated into an inflation forecast framework based on the p-star approach proposed by Czudaj (2011) among others.

¹⁹This question has already been tackled in the course of a correlation study by Spierdijk and Umar (2010) which is based on a standard VAR framework. As hedging measures Spierdijk and Umar (2010) apply the Pearson correlation coefficient, the Fisher coefficient in the Fama and Schwert (1977) regression, the hedge ratio proposed by Schotman and Schweizer (2000), the hedging capacity suggested by Bodie (1976) as well as the associated cost of hedging. Their analysis indicates that energy commodities and non-precious metals provide the best inflation hedges.

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Tables

Table 1: Johansen cointegration test (CPI)

Country	Lags	Hypothesis	Trace stat.	Cointegrating vector	(1, -1) χ^2
USA ^a	2	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	62.133**	$p_t - 0.408^{**}g_t - 0.000t$ (-5.594) (-0.532)	41.090** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	13.548*		
USA ^b	2	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	51.025**	$p_t - 0.433^{**}g_t$ (-8.757)	41.053** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	2.478		
UK ^a	4	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	84.096**	$p_t - 0.359^{**}g_t - 0.001^{**}t$ (-7.648) (-2.729)	28.105** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	11.323		
UK ^b	4	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	49.252**	$p_t - 0.455^{**}g_t$ (-10.377)	66.393** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	2.011		
Japan ^a	2	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	129.233**	$p_t - 0.231^{**}g_t - 0.001^{**}t$ (-3.570) (-4.934)	17.016** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	7.636		
Japan ^b	2	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	112.270**	$p_t - 0.177^{**}g_t$ (-4.290)	100.656** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	1.107		
Euro Area ^a	2	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	114.175**	$p_t + 0.006g_t - 0.002^{**}t$ (0.329) (-22.099)	48.848** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	7.322		
Euro Area ^b	4	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	36.976**	$p_t - 0.196^{**}g_t$ (-2.610)	25.477** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	2.589		

Note: *a*: Restricted trend, *b*: Unrestricted constant. * Statistical significance at the 5% level, ** at the 1% level. The *t*-statistics of the cointegration coefficients are given in parenthesis and the *p*-values of the LR test of the restricted model are reported in brackets. In case of insignificance the trend is deleted from the cointegrating space, allowing for a linear trend in the data, but not in the cointegrating equation. Critical values for testing (i) $H_0 : r = 0$ and (ii) $H_0 : r \leq 1$ are taken from MacKinnon et al. (1999): (a) 5% (i) 25.731 and (ii) 12.448, 1% (i) 31.153 and (ii) 16.553, (b) 5% (i) 15.408 and (ii) 3.841, 1% (i) 19.937 and (ii) 6.634, respectively.

Table 2: Johansen cointegration test (PPI)

Country	Lags	Hypothesis	Trace stat.	Cointegrating vector	$(1, -1) \chi^2$
USA ^a	2	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	27.134*	$p_t - 0.230^{**}g_t - 0.002^{***}t$ (-6.841) (-10.171)	16.496** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	5.315		
USA ^b	2	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	7.726		
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	1.731		
UK ^a	3	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	74.223**	$p_t - 0.390^{**}g_t - 0.001t$ (-7.211) (-1.460)	21.011** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	11.454		
UK ^b	3	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	63.339**	$p_t - 0.432^{**}g_t$ (-10.647)	59.933** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	1.523		
Japan ^a	3	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	30.943*	$p_t - 0.158^{**}g_t + 0.000t$ (-3.221) (0.394)	22.531** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	6.298		
Japan ^b	3	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	25.671**	$p_t - 0.152^{**}g_t$ (-3.358)	25.633** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	1.139		
Euro Area ^a	4	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	26.178*	$p_t + 0.055^{**}g_t - 0.002^{***}t$ (-2.452) (-21.585)	14.895** [0.000]
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	1.139		
Euro Area ^b	3	$H_0 : r = 0$ vs. $H_1 : r \geq 1$	11.696		
		$H_0 : r \leq 0$ vs. $H_1 : r \geq 2$	2.852		

Note: See Table 1.

Table 3: MS-VECM (CPI)

Country	Adjustment coefficients				Transition probability matrix P	
	Regime 1		Regime 2			
	Δp_t	Δg_t	Δp_t	Δg_t		
USA ^b	-0.002*	0.007	-0.005	-0.006	$p_{11} = 0.947^{**}$ (43.050)	$p_{21} = 0.053^{**}$ (2.410)
	(-2.000)	(0.780)	(-1.670)	(-0.170)	$p_{12} = 0.117^{**}$ (3.000)	$p_{22} = 0.883^{**}$ (22.640)
UK ^a	-0.007**	-0.011	-0.015**	-0.047	$p_{11} = 0.942^{**}$ (58.880)	$p_{21} = 0.058^{**}$ (3.630)
	(-7.000)	(-1.100)	(-3.750)	(-1.210)	$p_{12} = 0.237^{**}$ (3.820)	$p_{22} = 0.763^{**}$ (12.310)
Japan ^a	-0.011**	-0.045	-0.002*	0.010	$p_{11} = 0.734^{**}$ (7.890)	$p_{21} = 0.266^{**}$ (2.860)
	(-3.670)	(-1.550)	(-2.000)	(1.110)	$p_{12} = 0.093^{**}$ (3.100)	$p_{22} = 0.907^{**}$ (30.230)
Euro Area ^b	-0.002*	0.012	-0.006**	-0.006	$p_{11} = 0.960^{**}$ (53.330)	$p_{21} = 0.040^*$ (2.220)
	(-2.000)	(0.710)	(-6.000)	(-0.220)	$p_{12} = 0.068^{**}$ (2.430)	$p_{22} = 0.932^{**}$ (33.290)

Note: *a*: Restricted trend, *b*: Unrestricted constant. * Statistical significance at the 5% level, ** at the 1% level. *t*-statistics are given in parenthesis. The intercepts, the autoregressive coefficients and the variance-covariance matrix are not shown to save space.

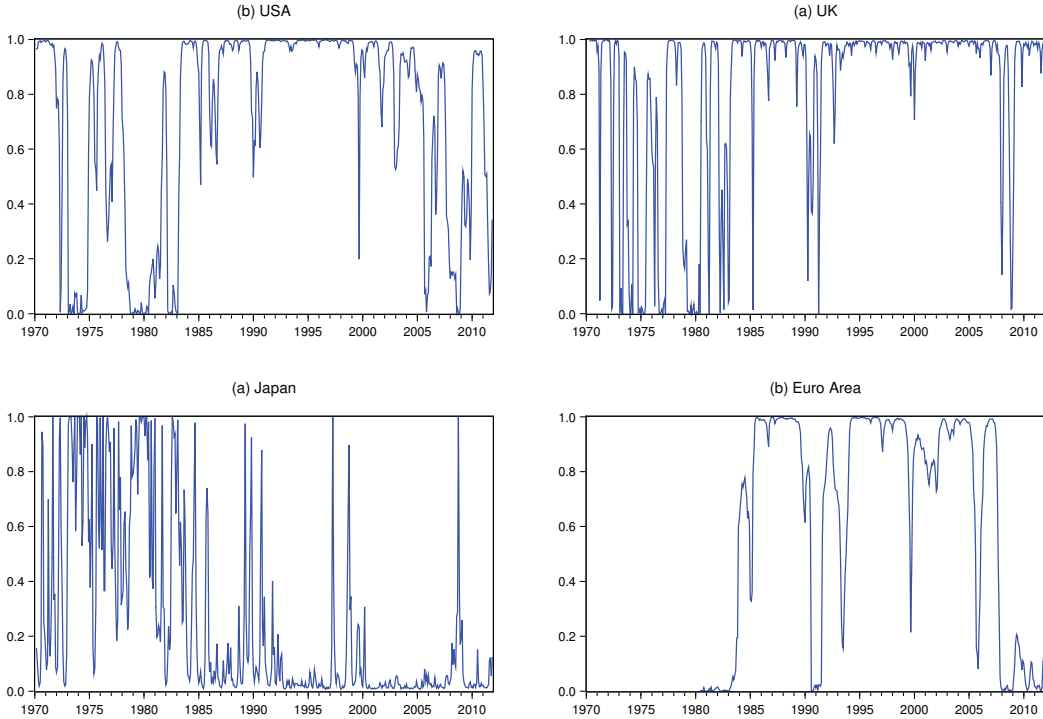
Table 4: MS-VECM (PPI)

Country	Adjustment coefficients				Transition probability matrix P	
	Regime 1		Regime 2			
	Δp_t	Δg_t	Δp_t	Δg_t		
USA ^a	-0.023	-0.032	-0.016**	-0.059	$p_{11} = 0.842^{**}$ (19.140)	$p_{21} = 0.158^{**}$ (3.590)
	(-1.770)	(-0.400)	(-4.000)	(-1.740)	$p_{12} = 0.072^{**}$ (3.600)	$p_{22} = 0.928^{**}$ (46.400)
UK ^b	-0.004**	-0.005	-0.006**	-0.025	$p_{11} = 0.962^{**}$ (64.130)	$p_{21} = 0.038^{**}$ (2.530)
	(-4.000)	(-0.500)	(-3.000)	(-0.860)	$p_{12} = 0.111^{**}$ (2.780)	$p_{22} = 0.889^{**}$ (22.230)
Japan ^b	-0.015	0.020	-0.002*	-0.047*	$p_{11} = 0.810^{**}$ (14.210)	$p_{21} = 0.190^{**}$ (3.330)
	(-1.880)	(0.220)	(-2.000)	(-2.140)	$p_{12} = 0.041^{**}$ (2.730)	$p_{22} = 0.959^{**}$ (63.930)
Euro Area ^a	-0.010**	-0.106	-0.001	-0.102	$p_{11} = 0.833^{**}$ (20.830)	$p_{21} = 0.167^*$ (4.180)
	(-3.330)	(-1.610)	(-0.130)	(-1.020)	$p_{12} = 0.269^{**}$ (3.790)	$p_{22} = 0.731^{**}$ (10.300)

Note: See Table 3.

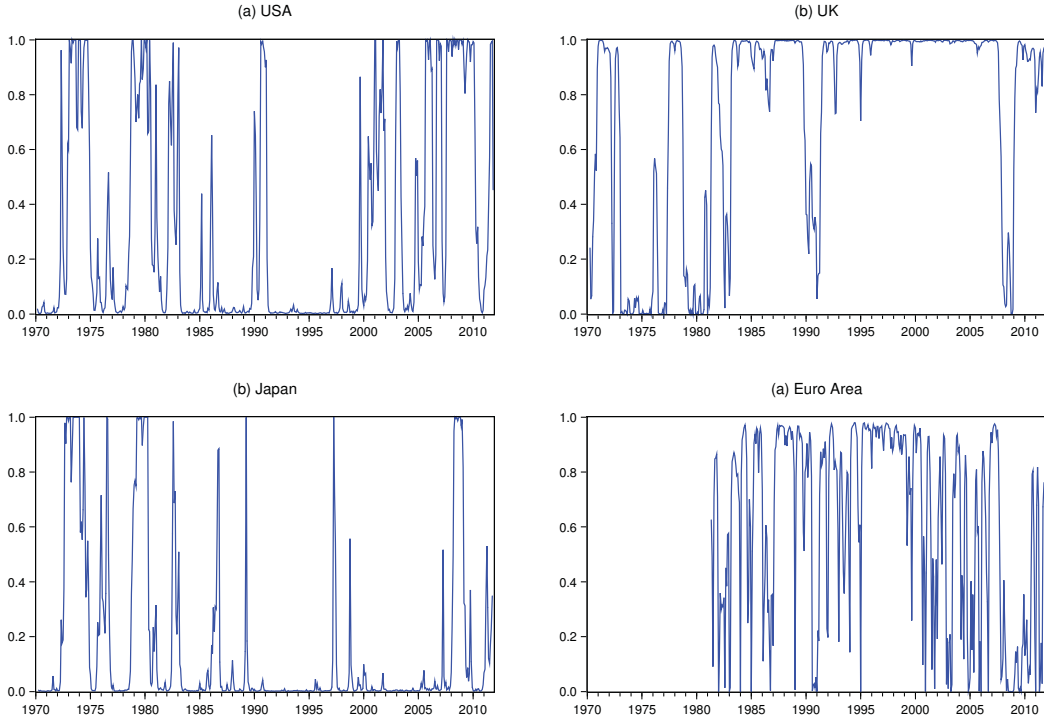
Figures

Figure 1: Smoothed probabilities for regime 1 for each country (CPI)



Note: (a) Restricted trend, (b) Unrestricted constant. The *RCM* for the USA, the UK, Japan, and the Euro Area is 29.6193, 18.5831, 31.4607, and 26.2708, respectively.

Figure 2: Smoothed probabilities for regime 1 for each country (PPI)



Note: (a) Restricted trend, (b) Unrestricted constant. The *RCM* for the USA, the UK, Japan, and the Euro Area is 24.8850, 19.0028, 16.1168 and 45.3926, respectively.